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Unemployment Duration in a Non-Stationary Macroeconomic Environment

By
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28 May, 1999

Unemployment Duration in a Non-Stationary Macroeconomic Environment

By Knut Røed and Tao Zhang*

Abstract

We investigate how unemployment exit probabilities are affected by economic incentives, spell duration and macroeconomic conditions. Building on a database containing *all registered unemployment spells in Norway in 1989-1998*, we apply an econometric model in which exit probabilities vary freely over spell durations as well as calendar time. We find that i) the replacement ratio affects the exit rate negatively and that this effect is counter-cyclical and stable over spell durations; ii) the exit rate rises just prior to benefit exhaustion; and iii) the predicted exit rate declines as the spell lengthens, but this may be explained by unobserved heterogeneity.

JEL Classification: C41, J64.

Keywords: Unemployment spells, business cycles, unemployment income, semi-parametric duration analysis.

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1 Introduction

This paper investigates how the probability of exiting unemployment is affected by economic incentives, unemployment duration, macroeconomic conditions and various individual characteristics. We apply an econometric model tailored for non-stationary macroeconomic environments in which individual exit probabilities vary both over spell durations and calendar time. With the aid of Norwegian register data, we combine the event histories of several different unemployment cohorts in order to disentangle duration- and calendar time effects.

The determinants of individual unemployment durations have received considerable attention in the literature (see e.g. Devine and Kiefer, 1991, or Pedersen and Westergård-Nielsen, 1998, for recent surveys); and thanks to the relatively easy access to large bases of administrative unemployment registers, a number of recent contributions have taken advantage of Norwegian micro data (Hernæs and Strøm, 1996; Bratberg and Vaage, 1996; Berg and Børing, 1997; Børing, 1998; Røed et al, 1999; Røed and Zhang, 1999a). These, as well as most of the duration studies from other countries, have relied on the analysis of single (or few) cohorts of unemployed persons. In cohort studies, elapsed duration and calendar time are inseparable. Hence, the development of exit probabilities over time cannot, without additional information, be assumed to pin down individual duration dependence (or unobserved heterogeneity), it may as well be attributed to aggregate seasonal- or business cycles. A popular method for identifying the unemployment duration pattern is to condition exit probabilities on the contemporary (local or national) aggregate rate of unemployment (Dynarski and Sheffrin, 1990; Meyer, 1990; Narendranathan and Stewart, 1993; Bratberg and Vaage, 1996; Rosholm 1996; Thoursie, 1998). However, the rate of unemployment is not only determined by the present state of labour demand, but also by previous states as well as by the present composition of the unemployment pool. Hence, a given rate of unemployment may have different implications for the exit probability at different phases of the business cycle. An alternative method for discriminating between duration- and calendar time effects, devised by Imbens and Lynch (1993) and Imbens (1994), is to switch the conventional roles played by duration and calendar time in the otherwise standard partial-likelihood procedure, such that calendar time effects are left unconstrained, while duration effects are parameterised.

In this paper, we identify the duration- and calendar specific components of the baseline hazard without parametric restrictions on either of the two components, by employing data for 100 monthly cohorts of entrants into the unemployment pool in Norway during the 1989-1998 period. We apply a discrete hazard rate model incorporating unconstrained calendar time- and duration effects, as well as person-specific time-varying covariates. The model is used to re-examine a number of issues that have been at the forefront of unemployment duration analysis, including the direction (and degree) of duration dependence and the effect of unemployment benefits. A particular emphasis is directed towards the interdependence between individual unemployment duration and the aggregate state of the labour market.

The next section gives a brief description of the data we use. Section 3 describes the potential pitfalls associated with volatile aggregate outflow rates. Section 4 presents the econometric model. Section 5 contains the econometric analysis and a discussion of the results in relation to previous findings. Section 6 concludes.

2 The Data

The data set comprises event histories of all workers who became full-time unemployed in Norway during the period from January 1989 to March 1998. In order to be sure that the persons we look at are really entrants into the unemployment pool, we define entrance in month *t* as being unemployed that month, but not in any of the previous six months. We observe the subsequent unemployment records on a monthly basis, until March 1998. The data display a pattern of frequent movements into and out of the unemployment register. This is partly related to slack registration behaviour, and partly to the fact that many people do manage to escape unemployment from time to time through temporary jobs or other activities (Røed et al, 1999). We assume that register-dropouts of only one month are erroneous and reconstruct the appropriate unemployment status for these secluded months. Exits from the register lasting more than one, but less than six months are considered ‘temporary’, while exits lasting at least six months are interpreted as ‘final’. In this paper, we focus on final exits. A person makes a final exit in month *t* if he is registered as unemployed in month *t-1*, is not registered in month *t*, and does not return to the register within the next five months (*t+1*,*t+2*,...,*t+5*). The duration of the unemployment spell is then calculated as the total number of months spent in the register from entry to final exit. This implies the 'clock is stopped' during a temporary exit. As we require six months of observations both to identify entries and exits, the effective observation period is reduced to July 1989-October 1997 (100 months). This leaves us with approximately 1.7 million unemployment spells with 13.5 million monthly unemployment observations.

The data set contains information about a number of individual characteristics, such as age, gender, marital status, children, educational attainment, county of residence, previous income, work experience, and, for immigrants, country of origin. The income- and work experience covariates are calculated on the basis of pensionable income records administered by The National Insurance Service. These records contain information about accumulation of pension points for all Norwegian citizens, starting from 1967. As a measure of work experience, we use the number of years with positive point accumulation out of the last 22 years; and as a proxy for earnings capacity, we use the average level of that accumulation (zero for persons with no work experience). The covariates also include a number of (time-varying) variables intended to capture unemployment status, e.g. current or previous labour market program experience, the amount of unemployment benefits or program participation payments, and the replacement ratio.

The replacement ratio is the payment received from the labour office, either as unemployment benefit or as program participation payment, relative to expected income when employed. We do not have direct information about unemployment benefits actually received, but we have the information required to calculate each person's benefit entitlement. For persons with a previous employment history, we assume that expected income is equal to previous income, corrected for average nominal wage growth. Persons without a previous employment history are not entitled to unemployment benefits; hence many of them have a zero replacement ratio. During participation in labour market programs however, they receive payments from the labour office. For some types of employment programs, these payments are set as a particular fraction of the going wage rate applying for the particular type of work, and in these cases we use this fraction as the replacement ratio. In these cases, we have knowledge about the replacement ratio, but not the level of the payment. For other types of programs we use a more 'rule-of-thumb' approach. For details about Norwegian unemployment benefit system and the construction of the replacement ratio, see Røed and Zhang (1999b).

3 The Calendar Time Pattern

The labour market environment faced by a cohort of unemployed workers typically changes considerably over time. First, there is a substantial cyclical, as well seasonal, variation in aggregate exit rates. Secondly, job starts are not spread out randomly over the calendar, but lumped at certain customary dates. Finally, unemployed persons are in reality never tracked continuously; instead, they are observed at certain points in time (in our case by the end of each month), and these points in time are sometimes affected by holidays (implying for example low outflow rates in December). If these calendar time factors are disregarded, and hence all observed time variation is interpreted as arising from duration dependence or unobserved heterogeneity, the resultant estimates will be strongly affected by the calendar time at which the cohort starts out.

The volatility of aggregate outflow rates in Norway is illustrated in Figure 1. The seasonal pattern is highly visible. A simple ratio-to-moving-averages seasonal decomposition method (Makridakis et al, 1983) indicates large month-to-month variations, with outflow rates during the 'best' summer months (August, September) almost twice as high as in the 'worst' winter months (November, December). For some reason, all the previous studies based on Norwegian register data that we know of (Hernæs and Strøm, 1996; Bratberg and Vaage, 1996; Berg and Børing, 1997, Børing, 1998; Røed et al, 1999) have analysed Autumn-entrants, and all but one of them (Bratberg and Vaage, 1996) have analysed October-entrants. As October-entrants typically face very low outflow rates during the first months of their spell, it is not difficult to imagine that attempts to characterise the degree of duration dependence may have been seriously distorted. In addition to that, there is an element of business cycle movements in the outflow rate. No matter the particular choice of unemployment-cohort in Norway the past 10 years, the cohort is bound to start either on a path of cyclically declining or cyclically rising aggregate outflow rates.

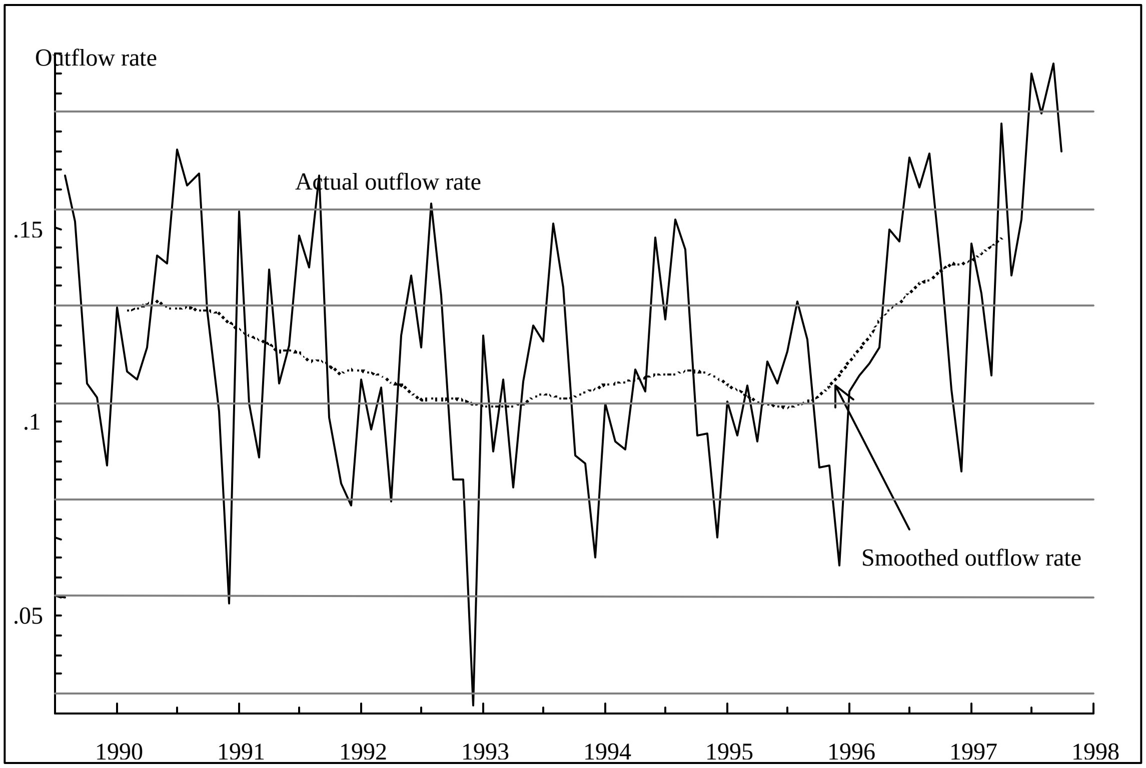

Figure 1. Outflow rates from unemployment in Norway August 1989-October 1997.
Note: Outflow rates are calculated on the basis of the whole unemployment pool in Norway (not only the cohorts described in section 2). Outflow in month *t* is defined as being unemployed in month *t-1*, but not in month *t* or any of the next five months (t+1,...,t+5). The smoothed series is a 13 months moving average (with half weights attached to the two end points).

4 The Econometric Model

In order to disentangle calendar time and duration effects, we build on a semi-proportional hazard rate framework. Since we observe unemployment status by the end of each month only, we develop the model in terms of discrete hazard rates. The discrete hazard rate for an individual i in month t, i.e. the probability of making a final exit during month t, conditioned on being unemployed at the start of the month, is defined as:

$$h_{it} = 1 - \exp\left(- \int_{t-1}^{t} q_i(u)du \right), \qquad (1)$$

where q_i is the underlying continues time hazard rate. Assume that for some group j of unemployed persons, the hazard rate can be factorised into a person-specific part that depends on calendar time, and a group-specific duration baseline that does not depend on calendar time. Let x_{it} be a vector of relevant individual characteristics, let t_{it} be individual unemployment duration, let s_{jt} be a measure of the 'state of the labour market' (labour demand) that faces group j in calendar month t, and let $b_j(t)$ be the group-specific continues time duration-baseline. The integral in (1) can then be rewritten as $\int_{t-1}^{t} q_i(u)du = f_{jt}(x_{it}, s_{jt}) \int_{t_{it-1}}^{t_{it}} b_j(u)du$, where $f_{jt}(.)$ is a non-negative function and $(x_{it} s_{jt})$ is (for simplicity) assumed constant within each month. Unemployment duration is of course closely related to calendar time. Let t_{i0} be the calendar time of entry and let o_{it} be the accumulated number of months spent in temporary exit. We then have that $t_{it}=t-o_{it}-t_{i0}$. Define $l_{jt} = \log\left(\int_{t-1}^{t} b_j(u)du\right)$ and let $f_{jt}(x_{it}, s_{jt}) = \exp(x_{it}'b_{jt} + s_{jt})$. We then have that the monthly exit probabilities to be estimated are given by

$$
\begin{gathered}
h_{jitt} = 1 - \exp\left(-\exp(x_{it}'b_{jt} + s_{jt} + l_{jt})\right), \\
j = 1,2,\ldots,J, \quad i = 1,2,\ldots,N_j, \quad t = 0,1,\ldots,T, \quad t = 1,2,\ldots,T.
\end{gathered} \tag{2}
$$

Parametric restrictions on the continuous time duration-baseline may be imposed by laying the appropriate restrictions on the l_{jt}-parameters. For example, Weibull-distributed durations would entail that $l_{jt} = \log(t^{a_j} - (t-1)^{a_j})$, where a_j is the single duration dependence parameter (less than unity in the case of negative duration dependence). Note that $t=0$ is the entry month for the first cohort (July 1989) and $t=T$ is the time at which still ongoing durations are censored[1] (October 1997). The first cohort is potentially observed for T months, the second for $T-1$ months and so on until the last cohort, which is observed for only one month. If we impose the restriction that $b_{jt}=b_j$ (constant parameters), we obtain a mixed proportional hazard model, i.e. a hazard rate that can be factorised into three components, a duration baseline, a calendar time baseline and a factor of proportionality that depends on individual characteristics only. The proportionality assumptions may be relaxed through the inclusion of interaction terms between individual characteristics, calendar time and duration.

It is obvious that the calendar- and duration parameters in equation (2) cannot be identified on the basis of a single cohort of unemployed persons without transitory exits, even in the pure proportional model. With only one cohort, the flexible baseline hazard typically estimated (Meyer, 1990; Narendranathan and Stewart, 1993; Arulampalam and Stewart, 1995) looks like

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[1] We also censor observations just prior to the retirement age of 67 years.

$$h_{jit} = 1 - \exp\left(-\exp(x_{it}' b_j + g_{jt})\right). \tag{3}$$

It is clear that the coefficients in g_{jt}, although often interpreted as duration effects, are really mixtures of the true duration-baseline coefficients and the aggregate state of the labour market, i.e. $g_{jt} = l_{jt} + s_{j\,t_0+t}$. The inclusion of e.g. the rate of unemployment in x_{it} as a time-varying covariate may attenuate, but hardly remove the labour demand effect[2].

The existence of unobserved heterogeneity imposes a negative bias in estimates of duration dependence and may also introduce bias in, or rather change the interpretation of, the estimates of other coefficients[3]. A popular remedy is to impose a particular distribution for unobserved heterogeneity (typically a Gamma distribution or a limited number of mass points), and then to estimate the parameters of this distribution together with the parameters of interest. However, this procedure requires some form of restrictions on the duration-baseline. If one has no additional information that can justify the restrictions, the resulting identification is questionable. Narendranathan and Stewart (1993, p. 71) argue that one may create distortions that are equally serious as those resulting from ignoring unobserved heterogeneity. It is also possible to identify the degree of true duration dependence conditional on the validity of a mixed proportional hazard rate model (Van den Berg and van Ours, 1996; Abbring et al, 1999). However, this condition does not have any theoretical foundation either (and in our case it is strongly rejected by the data). Fortunately, the bias arising from omitted heterogeneity is smaller the more flexible is the duration-baseline and the more explanatory power is embedded in observed covariates. In this paper, we primarily pursue what Blossfeld and Rohwer (1995, p. 256) consider the *only* empirically promising strategy to cope with unobserved heterogeneity: "To look for more and better data". In order to sort out the robust from the fragile results, we estimate a number of alternative duration models for a number of different groups of unemployed. As a part of this strategy, some models are also estimated with unobserved heterogeneity.

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[2] In our data, the current aggregate rate of unemployment accounts for less than 30 per cent of the variation in the current aggregate rate of outflow (the squared correlation coefficient is 0.29).

[3] For example a coefficient attached to a variable that captures educational attainment may partly measure the effect of education and partly the degree to which people with high education are more motivated than others.

5 Econometric Analysis and Results

In this section, we estimate a number of alternative formulations of equation (2), and discuss the results in relation to previous findings, both in Norway and in other countries. The models differ in terms of the group composition (j) and in terms of parametric restrictions (p). A brief description of the models we estimate is given in Table 1. The groups are constructed along the following dimensions: Previous labour market status (full-time job or education), type of unemployment (recall or not recall), gender and age. A particular emphasis is devoted to the hazard rates for prime aged, and previously full time employed persons. Members of this group (particularly males) are not likely to leave the labour force; hence the estimated hazards primarily reflect transitions to jobs.

The explanatory variables can be divided into three categories, according to the way they are updated during the spells. The first category consists of variables that are updated every month. The replacement ratio belongs to this category, together with four labour market program participation dummies (current and previous participation on training course or employment program). The second category consists of variables that are updated yearly (in January). It includes age, marital status, children and county of residence. The final category consists of variables that are constant throughout a spell. These are previous work experience, previous income, educational attainment, country of origin (immigrant status), gender, and the calendar month of inflow. The coefficients attached to the explanatory variables (b_{jt}) are either restricted to remain constant throughout the estimation period or restricted to remain constant within each of four different time periods; i) August 1989-May 1991 ii) June 1991-August 1993, iii) September 1993-August 95, and iv) September 1995-October 1997. These periods correspond to the different phases of the business cycle, starting with a moderate downturn in the first period, continuing with a recession in the second, followed by a recovery in the third, and then closing with a boom in the last period.

Table 1a

The Grouping of the Data and Descriptive Statistics

Group (j)	All (A)	Prime Aged Men (PAM)	Prime Aged Women (PAW)	Old Men (OM)	Old Women (OW)	Young from Employm. (YE)	Young from School (YS)	On Recall (R)
Recall/not recall	Both	Not recall	Not recall	Not recall	Not recall	Not recall	Not recall	Recall
Age at start of spell	16-66	25-50	25-50	51-66	51-66	16-24	16-24	16-66
Gender	Both	Men	Women	Men	Women	Both	Both	Both
Situation before spell	Any	Full time employed	Full time employed	Full time employed	Full time employed	Full time employed	In education	Full time employed
# spells	448528	229036	163127	39748	22957	170786	224797	278390
# months	3466816	2059430	1334505	587371	273988	1292198	1615846	1277358
Means:								
Age	33.13	36.05	35.68	60.10	58.84	22.54	20.44	37.98
Work exp. last 22 years	7.92	12.70	9.60	21.09	16.26	2.35	0.44	12.80
# pension points	1.80	2.93	1.98	4.50	2.33	1.15	0.17	3.00
Repl. ratio	0.49	0.49	0.53	0.50	0.55	0.51	0.46	0.50
Fractions:								
Men	54.83	100.00	0.00	100.00	0.00	56.90	47.75	78.26
< 11 years education	51.95	52.15	53.65	62.37	81.21	49.43	45.88	59.15
11-12 years education	37.77	35.66	33.39	24.44	12.63	46.76	47.32	35.72
> 12 years education	10.28	11.99	12.96	13.19	6.15	3.81	6.79	5.13
Immigrants from non-OECD	7.19	8.22	5.66	1.68	1.26	3.39	7.55	2.87
Training course	10.54	10.04	12.55	2.58	3.48	10.59	10.74	4.47
Employm. program	17.37	10.71	12.00	6.98	6.97	21.20	42.65	5.70
Benefits	57.12	70.26	68.93	84.48	82.60	56.52	16.10	79.73

Note: Group A is a random sample. Since most covariates are time varying, means and fractions are calculated with the total number of unemployment months as the base.

Table 1b

The Estimated Models and their Parametric Restrictions

Model type (p)	1	2	3	4	5	6
Duration baseline	Flexible	Flexible	Flexible	Weibull	Weibull	Weibull
Calendar baseline	Unrestricted	Unrestricted	Unrestricted	Unrestricted	Restricted	Restricted
Time varying parameters	Yes	Yes	No	Yes	No	No
Benefit countdown dummies	No	Yes	Yes	Yes	No	No
Age function	Quadratic/ Flexible	Quadratic	Unrestricted	Quadratic	Unrestricted	Unrestricted
Work experience function	Linear	Linear	Piecewise constant	Linear	Piecewise constant	Piecewise constant
Unobserved heterogeneity	No	No	No	No	No	Gamma

Note: The models and their motivation are more thoroughly explained in the text.

There are no restrictions on the calendar-time baselines. The group-specific state-of-the-market parameters (s_{jt}) are estimated freely for each calendar month and for each model. The duration baselines are also estimated without any restrictions at all for the first two years of unemployment spells. During the next three years, we assume that the baseline is constant within each three-month duration interval. After five years of unemployment, we assume that the duration baseline remains constant.

The models are estimated with a Maximum Likelihood method based on the Newton algorithm implemented in the TDA computer program (see Rohwer, 1998, for details). Let E_j denote the set of uncensored observations in group j. The log-likelihood function associated with group j can then be written as

$$\ell_j = \sum_{i \in E_j} \log\left(h_{jitt_i}\right) + \sum_{i \in N_j} \sum_{l=1}^{t_i^*} \log\left(1 - h_{jitl_i}\right), \tag{4}$$

where h_{jitt_i} is defined in (2), $t_i^* = (t_i - 1)$ for $i \in E_j$ and $t_i^* = t_i$ for $i \notin E_j$.

Our presentation of the results is organised in the following way: We first describe the estimated calendar time- (section 5.1) and duration (section 5.2) baseline hazard rates. We then present estimates concerning the replacement ratio (section 5.3) and participation in labour market programs (section. 5.4). Finally, we describe the estimated effects of education and labour market experience (section 5.5), demographic factors (section 5.6), and nationality (section. 5.7). As the various estimates are obtained from a number of different models, we indicate for each set of estimates the particular model from which it is generated, using the notation in Table 1. For example estimates obtained for prime aged males based on the model of type 1 (in Table 1b) is referred to as (PAM1). We typically report b_j-estimates based on the most general model (type 1 or 2) only.

5.1 The State-of-the-Market Baseline

In order to give a condensed picture of how labour demand has developed during the estimation period, we build in this subsection on models that are proportional in calendar time, i.e. models without time-varying parameters attached to explanatory variables. We normalise the state-of-the-market baseline such that $s_{j1}=0$ (August 1989). The relative monthly exit probabilities (h_{jitt}/h_{ji1t}) are then approximately equal to $\exp(s_{jt})$. Figure 2 depicts $(\exp(\hat{s}_{jt}))$ for all unemployed; in the upper panel together

with a 95 per cent point-wise confidence interval, and in the lower panel together with the actual outflow rate. The estimated baseline tracks the actual outflow rate closely (the squared correlation coefficient is 0.83), suggesting that most of the variation in the aggregate outflow rate is explained by variation in macroeconomic conditions (labour demand), rather than by variation in the composition of those becoming unemployed. This is in accordance with previous findings for Denmark (Rosholm, 1996), but at odds with some US evidence (see e.g. Darby et al, 1986)

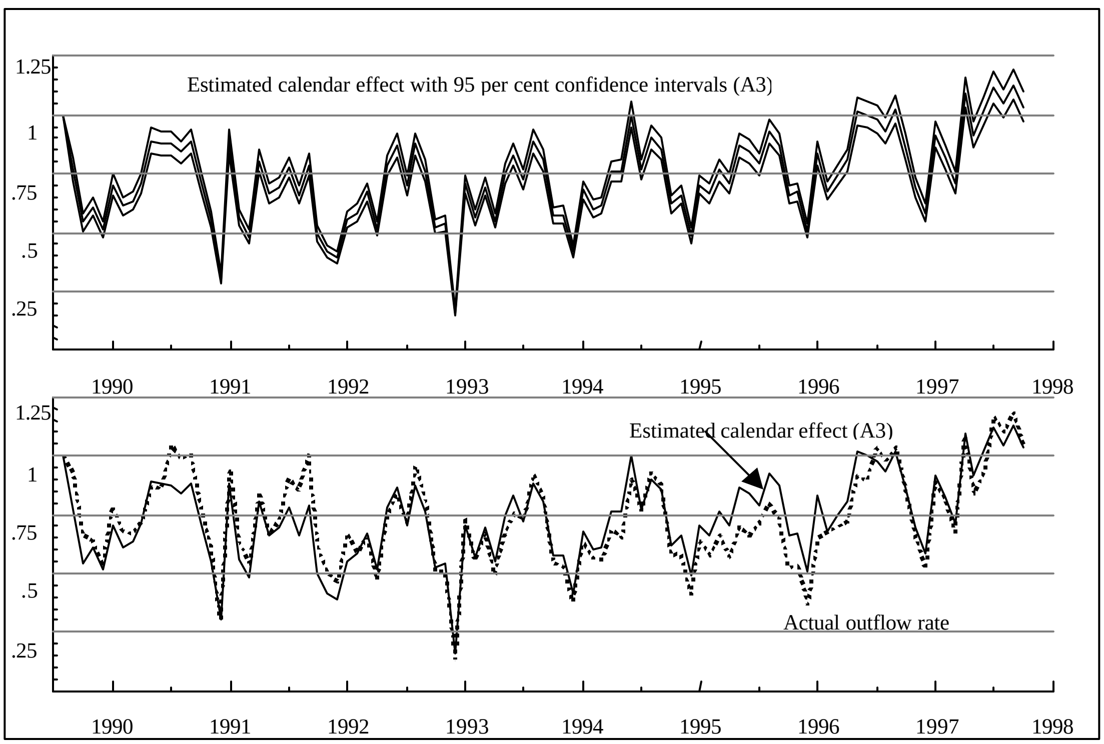

Figure 2. Estimated state-of-the-market baseline hazard rate (A3) with 95 per cent (point-wise) confidence intervals and actual aggregate outflow rates.
Note: Both series are normalised on the first month (i.e. they are equal to unity in August 1989)

Even though macroeconomic conditions do drive most of the fluctuations in outflow rates, composition effects also have a role to play. To focus more sharply on the business cycle pattern, Figure 3 plots smoothed calendar baseline estimates and outflow rates for prime aged men and women. These curves suggest that the macroeconomic conditions deteriorated sharply in 1990 and 1991, but that the subsequent recovery also started earlier than indicated by actual outflow rates. In fact, the estimates suggest that the trough was hit as early as in the autumn of 1991, despite that actual outflow rates continued to fall at least until the summer of 1993[4]. A corollary is that the unemployment pool at the start of the 1990's consisted of persons with relatively high individual exit rates. As we show in the next subsection, this is related to the duration-composition of the unemployment pool. The first part of a recession is typically characterised by a lot of short-term unemployment, but as the recession continuous, average spell duration increases. Either because of sorting mechanisms or because of true negative duration dependence, the composition of the unemployment pool deteriorates. This composition effect impedes the subsequent recovery; i.e. it takes time before a rise in individual exit rates (conditioned on elapsed duration) is transformed into improvements in actual outflow rates.

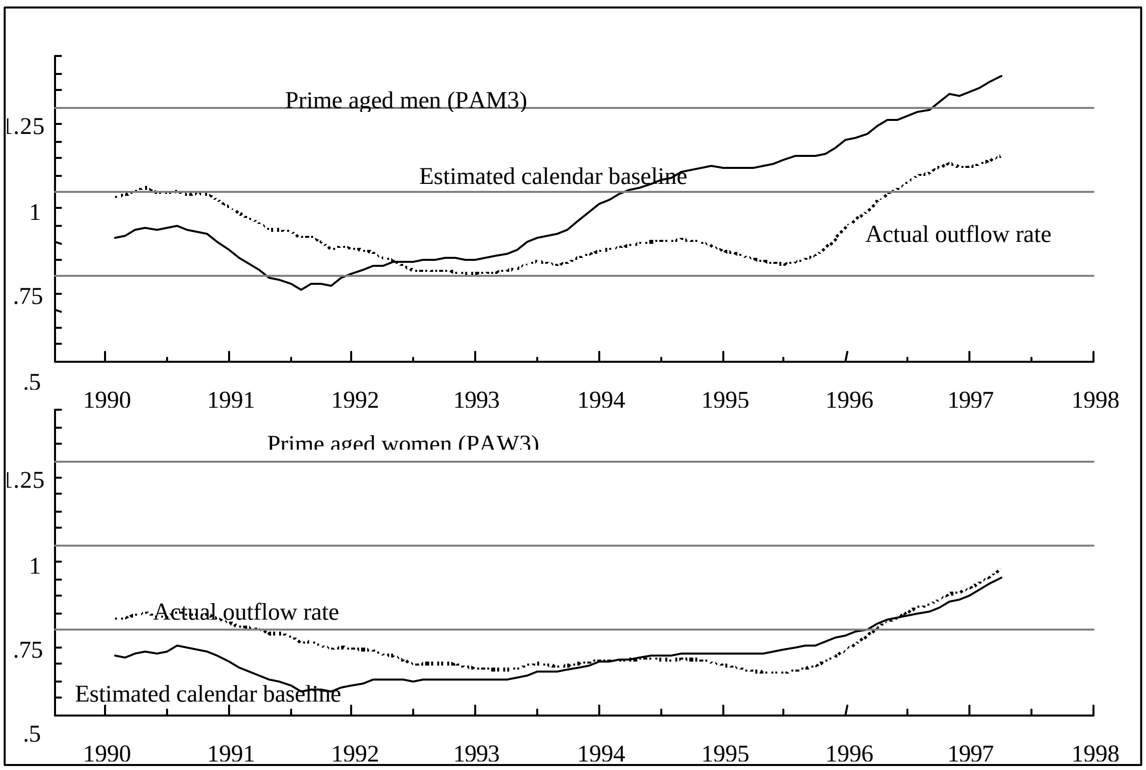

Figure 3. Smoothed estimates for the state-of-the-market baseline and actual aggregate outflow rates for previously full-time employed prime aged men and women.
Note: The smoothed series are 13 months centred moving averages of the original normalised (to August 1989) series (with half weights attached to the two end points).

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[4] This interpretation is supported by the fact that (the smoothed) *inflow rates* to unemployment also declined in late 1991 (see Røed and Zhang, 1999a), while labour force participation increased. Inflow rates rose slightly in late 1992, but that was related to the increase in the labour force participation rate.

5.2 The Duration Baseline

We normalise the duration baseline such that $l_{j1}=0$. The relative exit probabilities (h_{jitt} / h_{jit1}) are approximately equal to $\exp(l_{jt})$. The estimated duration baseline hazards $(\exp(\hat{l}_{jt}))$ for six different groups are depicted in Figure 4, together with 95 per cent (point-wise) confidence intervals.

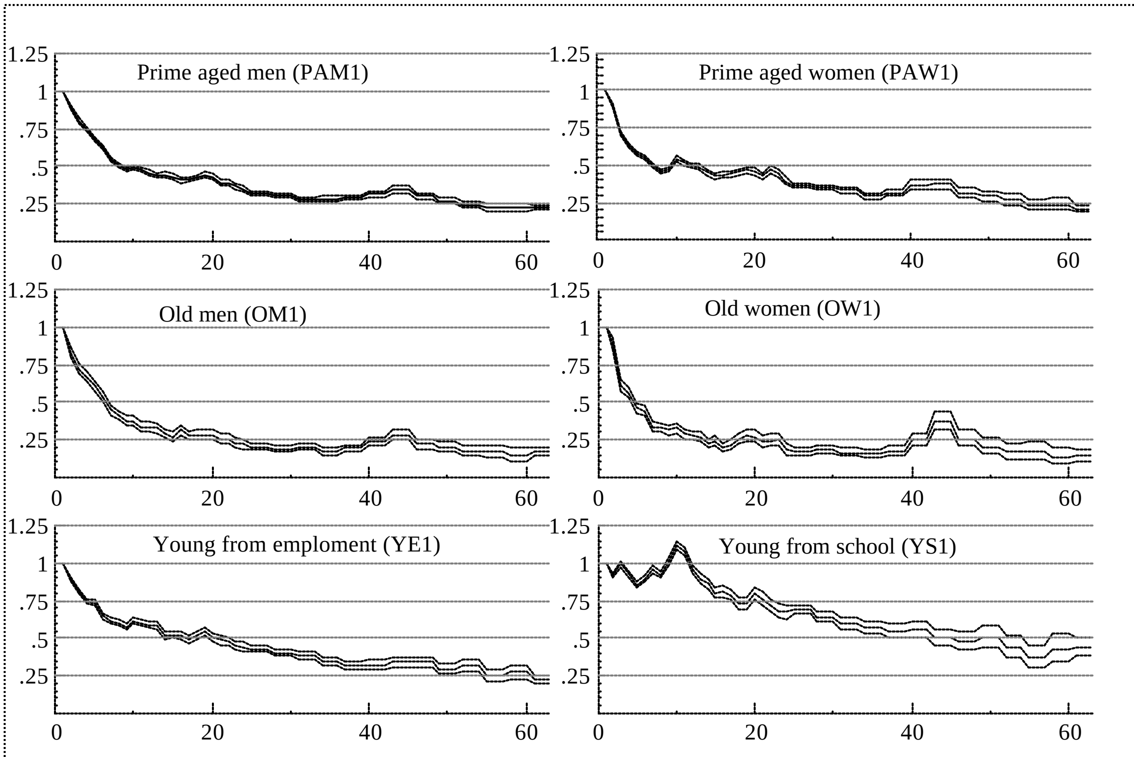

Figure 4. Estimated duration baseline hazard rates for ordinary unemployed (not on recall) measured in elapsed months (with 95 percent point-wise confidence intervals).

Note: The hazards are normalised to unity in first duration month. They are estimated with no parametric restrictions the first 24 months. The next 36 months, the hazards are assumed constant within each three-month interval. From duration month 62, the hazards are assumed constant.

The duration baselines - conditioned on observed covariates - slope sharply downwards for all groups except for young persons who entered unemployment directly from school. Negative duration dependence at the level of the individual may occur because long-lasting human inactivity is detrimental to search effort, health, and human capital (Phelps, 1972; Hargreaves Heap, 1980; Björklund and Eriksson, 1995), or because employers rank job applicants according to unemployment duration as a sort of statistical discrimination device (Blanchard and Diamond, 1994). But we cannot be sure that the duration dependence displayed in Figure 4 operates at the individual level. It may also result from the existence of unobserved heterogeneity, as those with the highest exit probabilities (conditioned on all observable information) tend to exit first. We return to this issue below.

Table 2
The Effect of Unemployment Benefit Exhaustion.

	Time Left to Temporary Benefit Exhaustion		
	3 months	2 months	1 month
Previously full time employed, not on recall:			
Prime aged men (PAM2)	0.0816	0.3464	0.2258
	(0.0577)	(0.0531)	(0.0604)
Prime aged women (PAW2)	0.0227	0.3325	0.3545
	(0.0732)	(0.0677)	(0.0719)
Old men (OM2)	0.2360	0.3313	0.6045
	(0.1754)	(0.1750)	(0.1634)
Old women (OW2)	0.2041	0.5238	0.6250
	(0.2237)	(0.1948)	(0.1974)
	Time Left to Permanent Benefit Exhaustion		
	3 months	2 months	1 month
Prime aged men (PAM2)	0.0555	0.1940	0.1138
	(0.0686)	(0.0687)	(0.0761)
Prime aged women (PAW2)	0.2502	0.1054	0.1768
	(0.0924)	(0.1076)	(0.1124)
Old men (OM2)	0.3634	0.3943	0.2568
	(0.1414)	(0.1500)	(0.1700)
Old women (OW2)	0.5241	0.7373	0.4328
	(0.1659)	(0.1610)	(0.1986)

Note: Standard errors in parentheses.
Likelihood ratio tests associated with the six countdown dummies yielded the following results:
Prime aged men (PAM2 vs. PAM1): $c^2(6)$=61.38;
Prime aged women (PAW2 vs. PAW1): $c^2(6)$=52.56
Old men (OM2 vs. OM1): $c^2(6)$=29.01
Old women (OW2 vs. OW1): $c^2(6)$=41.74
Young from employment (YE2 vs. YE1): $c^2(6)$=7.72

For all groups, there is a small rise in the hazards around the 18-22 duration months, and a somewhat larger rise around the 40-44 months. These periods correspond to the times at which unemployment benefits could be temporarily and permanently terminated[5] respectively, hence the results support to some extent previous findings that the hazard rate rises in anticipation of benefit exhaustion (see e.g. Meyer, 1990; Lindeboom and Theeuwes, 1993; Hunt, 1995; Carling et al, 1996; Thoursie, 1998). Table 2

[5] Until May 1992, benefit claimants were subject to a 13-week cut-off-period after 80 weeks of benefit exhaustion. After that, a new 80-week period could start. In the period from May 1992 to January 1997, an exemption rule applied, which effectively extended the benefit period to two consecutive periods of 93 weeks. Since 1997, there has been a formal limitation of 156 weeks, followed by some form of labour market program or (if such a program is not offered) more cash transfers.

displays the results from including benefit countdown dummies associated with the three last months just prior to exhaustion in the models[6]. Except for youths (who typically have very limited benefit entitlements), the countdown dummies do play a significant role. However, the effects are not very strong (and not very precisely determined) and it appears that a substantial part of the rise occurs just after (and not just prior to) exhaustion, suggesting that many of the exits at this stage are really exits out of the labour force. This is consistent with the previous finding by Bratberg and Vaage (1996) that the main effect of benefits running out is to make people drop out of the unemployment register.

Table 3
The Interaction Between Unemployment Duration and the State of the Labour Market

	Interaction effects with the estimated state of the labour market		
	Unemployed for at least 3 months	Unemployed for at least 6 months	Unemployed for at least 12 months
Previously full time employed, not on recall:			
Prime aged men (PAM2)	0.0345	0.1199	0.0639
	(0.0299)	(0.0336)	(0.0341)
Prime aged women (PAW2)	0.0127	0.0868	0.0323
	(0.0350)	(0.0393)	(0.0393)
Old men (OM2)	0.1230	-0.1702	-0.0904
	(0.0814)	(0.0899)	(0.0875)
Old women (OW2)	-0.0533	0.0667	-0.0420
	(0.0999)	(0.1155)	(0.1135)
Young men and women (YE1)	0.1528	0.1306	-0.1316
	(0.0352)	(0.0384)	(0.0398)
Previously full time employed, on recall:			
All (R1)	0.1041	0.1407	0.1977
	(0.0261)	(0.0368)	(0.0475)
Previously in education:			
Young men and women (YS1)	0.4733	-0.0690	-0.6080
	(0.0347)	(0.0352)	(0.0361)

Note: The interaction effects are additional, implying e.g. that the estimated effect associated with being unemployed for more than 12 months is the sum of the three coefficients. Standard errors in parentheses.

We also included an interaction term of long-term-unemployment dummies (unemployed for at least three, six or 12 months respectively) and the estimated state-of-the

[6] Note that these dummies do vary independently of unemployment duration for four reasons. First, during participation in labour market programs, benefits are not exhausted. Second, persons above 64 years of age are exempted. Third, some unemployed persons are not entitled to benefits at all. And finally, the first set of dummies (associated with the temporary exhaustion) applied only until May 1992. Note however that because of uncertainty with respect to the exact lengths of the spells, there is also some uncertainty with respect to the exact timing of benefit exhaustion.

labour market effects (from a first-stage estimation). The results are presented in Table 3. They reject the proportionality assumption, and indicate that long-term-unemployed are relatively more affected by business cycles than short term unemployed. This lends support to the ranking hypothesis proposed by Blanchard and Diamond (1994), i.e. that firms, when receiving job applications, hire the applicant with the shortest unemployment duration. During a recession, there are typically many applicants per vacancy; hence the ranking effect (and the associated negative duration dependence) becomes stronger. Our results are in accordance with previous findings by Dynarski and Sheffrin (1990) and Butler and McDonald (1986), but in conflict with results reported by Imbens and Lynch (1993) and Rosholm (1996). However, the positive interaction effects may also reflect unobserved heterogeneity and associated knock-on effects. When the labour demand is very low, highly qualified job seekers (in terms of unobservables) may be willing to accept jobs that under normal circumstances would have been available for the less qualified.

Apart from the rises associated with benefit exhaustion, the estimated duration baseline hazards display a fairly monotonic pattern. This is to some extent at odds with previous studies based on flexible baseline hazards. For example, Narendranathan and Stewart (1993), Kerckhoffs et al (1994), Arulampalam and Stewart (1995), Bratberg and Vaage (1996), Rosholm (1996) and Addison and Portugal (1998) all report hazards that embody substantial non-monotonicities. But, as these papers tell different stories about the shape of that non-monotonicity, it is difficult to derive any robust conclusions about the structural characteristics of the duration baseline. It may be hypothesised that the diverse results to some extent reflect the volatility in aggregate outflow rates, i.e. that they are driven by calendar time- rather than duration effects. Our estimated baseline hazards are in fact not very different from simple one-parameter monotonous Weibull hazards. This resemblance may serve as a platform for a conditional separation of individual duration dependence and unobserved heterogeneity. As discussed in the previous section, given that the Weibull model is a valid representation of individual hazards, we may apply this model, together with a representation of unobserved heterogeneity, in order to disentangle these two sources of negatively estimated duration dependence. We first estimated discrete Weibull hazards ($l_{jt} = \log(t^{a_j} - (t-1)^{a_j})$) for prime aged men and women[7], without unobserved heterogeneity. Although the Weibull restrictions were rejected on conventional statistical criteria, the deviations from the Weibull models were small (in quantitative terms). The estimated Weibull duration parameters (a_j) were estimated to 0.76 for men and 0.77 for women (with negligible standard errors). We then added to the Weibull model a Gamma distribution for unobserved heterogeneity. However, in order to solve the computational problems[8], we had to trim the model, both in terms of the number of observations and the number of explanatory variables. The results indicate that the duration dependence may be completely accounted for by unobserved heterogeneity. Non of the estimated duration parameters were significantly different from unity. An Exponential model, mixed with a Gamma distribution of unobserved heterogeneity, seems to fit the data well. More importantly, the coefficients other than those reflecting duration dependence turned out to be robust with respect to the selection of duration baseline model and the imposition of unobserved heterogeneity.

5.3 The Replacement Ratio

The estimated effects of the replacement ratio are presented in Table 4[9]. The estimates indicate a highly significant negative effect associated with higher replacement ratios. For example, for prime aged males, an increase in the replacement ratio with 10 percentage points decreases the hazard rate with approximately 4-10 per cent, depending on the period in question. Evaluated at the mean level of compensation this is compa-

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[7] To simplify the computations, we replaced $\log(t^{a} - (t-1)^{a})$ in the likelihood function with the linear approximation around $a=1$, i.e. $(t \log t - (t-1)\log(t-1))(a-1)$.

[8] The likelihood function for the mixture model is complicated by the fact that when person specific "fixed effects" are present, the unemployment spells can no longer be split into separate parts that depends on observed time-varying covariates only. We used a program written by Stephen Jenkins, University of Essex, to take account of this problem. It turned out to be impossible (at least with our computational resources) to estimate this model with the complete set of observations. The models reported here are based on approximately 40 000 spell observations each. The most important change in the model formulation is that the calendar time dummies are replaced by scalar variables containing point estimates for s_j from the Weibull model without heterogeneity.

[9] In the present paper, we focus on the economic incentives embedded in the (estimated level of) a single replacement ratio. Note that we cannot estimate an elasticity within this framework, as there are a number of zeros represented in the data. A more thorough investigation into the relationship between unemployment income and the hazard rate, involving alternative measures of unemployment income, expected employment income, as well as measures of benefit exhaustion, is in progress in Røed and Zhang (1999b).

rable to elasticities (with respect to the benefit level) ranging from –0.2 to –0.5. This indicates responses in line with previous findings for the United Kingdom (Narendranathan et al, 1985; Narendranathan and Stewart, 1993; Arulampalam and Stewart, 1995), but stronger responses than typically found in continental Europe (Hujer and Schneider, 1989; Groot, 1990; van den Berg, 1990; Steiner, 1990).

Table 4
The Estimated Effect of the Replacement Ratio

	Period 1	Period 2	Period 3	Period 4
Previously full time employed, not on recall:				
Prime aged men (PAM2)	-0.9800	-1.1403	-0.7863	-0.5027
	(0.0248)	(0.0199)	(0.0191)	(0.0181)
Prime aged women (PAW2)	-0.4263	-0.6479	-0.5329	-0.1160
	(0.0302)	(0.0243)	(0.0224)	(0.0197)
Old men (OM2)	-1.3197	-1.5065	-1.0652	-0.4268
	(0.0708)	(0.0581)	(0.0580)	(0.0509)
Old women (OW2)	-0.4406	-0.6457	-0.5319	0.2479
	(0.0925)	(0.0764)	(0.0691)	(0.0587)
Young men and women (YE1)	-0.5760	-0.8320	-0.6291	-0.3445
	(0.0249)	(0.0216)	(0.0216)	(0.0212)
Previously full time employed, on recall				
All (R1)	-0.4727	-0.7410	-0.7357	-0.3288
	(0.0158)	(0.0163)	(0.0183)	(0.0179)

Note: Standard errors in parentheses.

The results indicate that unemployment compensation reduces the hazard rate much more in recessions than in recoveries. The estimated effect is strongest in period 2 (the slump) and weakest in period 4 (the boom) for all groups. One interpretation of this result is that in periods with little labour demand, a lot of search effort is required in order to obtain a job, and hence individual economic incentives becomes more important. Our results at this point are contrary to findings for the United Kingdom (Arulampalam and Stewart, 1995) and the United States (Moffitt, 1985). Arulampalam and Stewart (1995) hypothesise that the negative benefit effect is pro-cyclical because demand constraints (the lack of job offers) are relatively more important compared to supply constraints (individual reservation wages) in recessions. But even though reservation wages do not play a prominent role during recessions, endogenously determined search intensity may be all the more important[10]. Moreover, the potential role of reservation wages during a boom may, for benefit receivers, be restrained by strict enforcement of job acceptance requirements. The finding of a coun-

[10] There is some evidence, based on the estimation of structural search models, indicating that reservations wages are empirically unimportant, and that virtually all job offers are accepted (van den Berg, 1990; Devine and Kiefer, 1991).

ter-cyclical benefit-effect may indicate that these requirements do work according to their intention; in periods with abundant labour demand, benefit receivers are compelled into the vacant jobs.

The results in Table 4 suggest that prime aged women are less sensitive towards the replacement ratio than men. But this is only true for the short-term unemployed. The models also contain interaction effects between the replacement ratio and unemployment duration. The results are given in Table 5. They indicate that, as the unemployment spell lengthens, the negative replacement ratio effect for women becomes stronger, while it remains constant for men. After three months of unemployment, the estimated effect for women is almost exactly the same as for men. Our results do not support the previous findings, reported by Nickell (1979), Fallick (1991), Narendranathan (1993) and Arulampalam and Stewart (1995), that unemployment income incentive effects are strongest in the beginning of the spell. But, as pointed out by Pedersen and Westergård-Nielsen (1998, p. 87), the few observations at high durations make these previous results rather tentative.

Table 5
Interaction of the Replacement Ratio and the Duration of Unemployment

	Interaction effects with replacement ratio		
	Unemployed for at least 3 months	Unemployed for at least 6 months	Unemployed for at least 12 months
Previously full time employed, not on recall:			
Prime aged men (PAM2)	-0.0262	-0.0179	0.0411
	(0.0216)	(0.0272)	(0.0282)
Prime aged women (PAW2)	-0.4780	0.0400	0.1176
	(0.0259)	(0.0336)	(0.0349)
Old men (OM2)	-0.0420	-0.1044	-0.2445
	(0.0613)	(0.0761)	(0.0778)
Old women (OW2)	-0.3744	-0.2758	-0.3799
	(0.0805)	(0.1036)	(0.1027)
Young men and women (YE1)	-0.1179	0.3458	-0.0760
	(0.0231)	(0.0287)	(0.0310)
Previously full time employed, on recall:			
All (R1)	0.0626	-0.5908	-0.0135
	(0.0203)	(0.0320)	(0.0424)

Note: The interaction effects are additional, implying e.g. that the estimated effect associated with being unemployed for more than 12 months is the sum of the three coefficients. Standard errors in parentheses.

Unemployed on recall (mostly men) are also less sensitive towards economic incentives. This is no surprise, since many of these workers are called back to their previous jobs without having to exert any search effort. However, even for this group, the low sensitivity only applies for the short-term unemployed. A likely explanation is that recall unemployment, after a while is transformed into ordinary unemployment, if it turns out that the previous employers do not have anything to offer after all.

5.4 Participation in Labour Market Programs

The models contain dummy variables that capture current, as well as previous participation in training courses and/or employment programs. Selected estimates are presented in Table 6. These estimates cannot be interpreted as pure treatment effects; they also reflect various selection mechanisms that are not accounted for in the model (self-selection, administrative selection). A relatively robust result is that participation in training courses is associated with a substantial decline in the hazard rate during the course. This effect appears to be much stronger in periods with relatively low unemployment (period 1 and 4) than in periods with high unemployment (period 2 and 3). This result probably reflects that, even though program participants are obliged to seek (and be available for) work during the course of a program, many participants do take a break in their search effort in order to complete the program first. For all groups, the lowest negative effect (in absolute terms) is estimated for period 2, which was a period with high and rapidly rising unemployment. Employment programs do not decrease the hazard during participation to the same extent as training courses. In fact, during the worst part of the slump, it appears that participation in employment programs raised the hazard rate significantly for adult men.

When the training course or the employment program is completed, the estimated hazard rate typically rises above its initial level for adults, but not for youths. For young persons, current or previous participation in any kind of labour market program is associated with a low predicted hazard rate. This could indicate that labour market programs are used as a sort of ‘storage’ for youths that do not fit into ordinary education (unemployed youths are automatically qualified for program participation in Norway). Most youths have not yet earned the entitlement to (a decent) unemployment benefit; hence program participation may be seen as the only way to earn some income. Previously full time employed adults on the other hand are likely to be more motivated, and to have been through a more thorough screening (both with respect to needs and potential returns) before they enter a program. Hence, the associated significant and positive effects on the hazard rate may reflect selection, as well as positive treatment effects. The result that previous program participation has strongest effects in periods with high labour demand suggests that treatment-effects do have a role to play (the general pattern is that heterogeneity is relatively more important in periods with low labour demand).

Table 6
Present and Previous Participation in Labour Market Programs

	Period 1	Period 2	Period 3	Period 4
Previously full time employed, not on recall:				
Prime aged men (PAM2)				
Current course	-0.3626	-0.0869	-0.3126	-0.5533
	(0.0355)	(0.0216)	(0.0193)	(0.0208)
Current employment program	-0.0600	0.2477	0.1158	-0.1350
	(0.0478)	(0.0267)	(0.0209)	(0.0221)
Previous course	0.1997	0.0432	0.1131	0.1871
	(0.0300)	(0.0164)	(0.0134)	(0.0137)
Previous employment program	0.1547	-0.0859	0.0398	0.0926
	(0.0420)	(0.0219)	(0.0165)	(0.0169)
Prime aged women (PAW2)				
Current course	-0.7813	-0.4482	-0.5698	-0.7378
	(0.0441)	(0.0266)	(0.0247)	(0.0230)
Current employment program	-0.4395	-0.1989	-0.1160	-0.4154
	(0.0608)	(0.0325)	(0.0257)	(0.0253)
Previous course	0.0597	0.0497	0.0756	0.1111
	(0.0370)	(0.0197)	(0.0169)	(0.0155)
Previous employment program	0.1155	0.0702	0.1139	0.1560
	(0.0539)	(0.0265)	(0.0205)	(0.0193)
Young men and women (YE1)				
Current course	-0.7395	-0.3956	-0.5222	-0.7226
	(0.0406)	(0.0250)	(0.0247)	(0.0264)
Current employment program	-0.3851	-0.0065	-0.1598	-0.5719
	(0.0313)	(0.0195)	(0.0201)	(0.0232)
Previous course	0.0309	-0.0174	-0.0110	0.0244
	(0.0319)	(0.0174)	(0.0164)	(0.0177)
Previous employment program	0.0296	-0.0396	0.0204	0.1225
	(0.0273)	(0.0161)	(0.0159)	(0.0175)
Previously in education:				
Young men and women (YS1)				
Current course	-0.8702	-0.2468	-0.4233	-0.6351
	(0.0436)	(0.0243)	(0.0222)	(0.0243)
Current employment program	-0.7972	-0.0338	-0.2726	-0.6307
	(0.0289)	(0.0175)	(0.0183)	(0.0221)
Previous course	-0.1505	-0.2582	-0.1159	-0.0341
	(0.0338)	(0.0172)	(0.0151)	(0.0160)
Previous employment program	0.0423	-0.0190	-0.0066	0.1124
	(0.0230)	(0.0141)	(0.0134)	(0.0145)

Note: Standard errors in parentheses.

5.5 Education and Work Experience

As indicated by the estimates reported in Table 7, there is a significant correlation between educational attainment and exit rates. Low education lowers- and high education raises the exit rates.

Table 7
The Estimated Effect of Education , Work Experience and Earnings Capacity

	Period 1	Period 2	Period 3	Period 4
Previously full time employed, not on recall:				
Prime aged men (PAM2)				
9 years education or less	-0.1559	-0.1666	-0.1515	-0.1116
	(0.0161)	(0.0118)	(0.0114)	(0.0119)
10 years education	-0.0775	-0.1630	-0.1650	-0.1262
	(0.0151)	(0.0104)	(0.0100)	(0.0105)
11-12 years education	0	0	0	0
	(ref.)	(ref.)	(ref.)	(ref.)
13-16 years education	0.0826	0.1079	0.0505	0.0218
	(0.0211)	(0.0147)	(0.0137)	(0.0123)
17 years education or more	0.1616	0.1977	0.1268	0.0798
	(0.0463)	(0.0348)	(0.0311)	(0.0237)
Work experience	0.0082	0.0199	0.0258	0.0104
	(0.0022)	(0.0017)	(0.0015)	(0.0013)
Average income in work career (approx.)	0.0599	0.0640	0.0762	0.0776
	(0.0059)	(0.0042)	(0.0040)	(0.0041)
Prime aged women (PAW2)				
9 years education or less	-0.1778	-0.1921	-0.1725	-0.1160
	(0.0220)	(0.0161)	(0.0153)	(0.0142)
10 years education	-0.0370	-0.1080	-0.1248	-0.0837
	(0.0192)	(0.0130)	(0.0124)	(0.0114)
11-12 years education	0	0	0	0
	(ref.)	(ref.)	(ref.)	(ref.)
13-16 years education	0.2328	0.1651	0.1395	0.1671
	(0.0257)	(0.0172)	(0.0154)	(0.0125)
17 years education or more	0.2994	0.2105	0.1084	0.1224
	(0.0677)	(0.0456)	(0.0406)	(0.0267)
Work experience	0.0074	0.0081	0.0098	0.0071
	(0.0020)	(0.0014)	(0.0014)	(0.0012)
Average income in work career (approx.)	-0.0347	-0.0112	0.0176	0.0149
	(0.0093)	(0.0065)	(0.0061)	(0.0055)
Previously in education:				
Young men and women (YS1)				
9 years education or less	0.0226	-0.0286	-0.0882	-0.0927
	(0.0160)	(0.0132)	(0.0116)	(0.0128)
10 years education	-0.1752	-0.0697	-0.2159	-0.1499
	(0.0150)	(0.0103)	(0.0106)	(0.0120)
11-12 years education	0	0	0	0
	(ref.)	(ref.)	(ref.)	(ref.)
13-16 years education	0.2802	0.3342	0.3043	0.2705
	(0.0203)	(0.0129)	(0.0148)	(0.0160)
17 years education or more	0.5340	0.4152	-0.0468	0.2274
	(0.1420)	(0.0914)	(0.1463)	(0.0644)
Work experience	0.0524	0.0599	0.0758	0.0810
	(0.0085)	(0.0053)	(0.0051)	(0.0063)
Average income in work career (approx.)	-0.0581	0.0125	0.0519	0.0348
	(0.0185)	(0.0120)	(0.0115)	(0.0133)

Note: Standard errors in parentheses.

Unsurprisingly perhaps, high education increases the hazard rate much more for individuals entering unemployment directly from school than for persons with more employment experience. The results in Table 7 indicate that persons with little education and work experience are relatively more sensitive to business cycles than highly skilled and experienced workers. Similar results have previously been reported for the Netherlands by Teulings (1993). The explanation is probably that when employment prospects are sufficiently meagre, educated persons are willing to accept jobs for which they are over-qualified. At the same time, employers take advantage of the excess labour supply to increase qualification standards for new hires. As a consequence, the competition for low-skill jobs becomes harder, and in this competition, the persons with lowest education yield. More surprisingly, perhaps, the results also indicate that educational attainment has become less important over time. For the prime aged, there is a clear pattern of convergence in education-specific hazard rates. This is at odds with the popular hypothesis that the European unemployment problem is driven by a relative deterioration of employment prospects for the low-skilled (OECD, 1994; Krugman, 1994). For young persons entering unemployment directly from school, on the other hand, the relative performances of persons with low education have deteriorated sharply. However, this is likely to reflect selection related to an educational reform in Norway that made 12 years education available for all. It is also likely that the apparently declining relative hazards for the very high-skilled school leavers reflect selection mechanisms, as enrolment into Universities and Colleges were expanded enormously during the recession.

5.6 Demographic Factors

Table 8 displays some estimation results regarding the effect of gender, age and family status. The difference in exit rates between single men and women was eliminated during the 1990's. However, marriage and childcare still affect men and women differently. Marriage is associated with higher exit rates for both sexes, but the effect is much stronger for men. Responsibility for small children implies that the estimated hazard rate for prime aged women falls with as much as 25-30 per cent, while the male hazard rate is left virtually unchanged.

Table 8
The Estimated Effect of Gender, Age and Family Status

	Period 1	Period 2	Period 3	Period 4
All (A2)				
Being a man	-0.0586	-0.0757	-0.0017	-0.0121
	(0.0100)	(0.0075)	(0.0073)	(0.0071)
Being married	0.1363	0.1801	0.1843	0.1520
	(0.0101)	(0.0079)	(0.0082)	(0.0080)
Having small children (up to 5 years)	-0.0208	-0.0573	-0.0704	-0.0524
	(0.0132)	(0.0102)	(0.0108)	(0.0120)
Having small children and being a women	-0.2482	-0.2105	-0.2647	-0.2941
	(0.0186)	(0.0141)	(0.0148)	(0.0157)
Age (deviation from mean)	-0.0101	-0.0201	-0.0243	-0.0223
	(0.0008)	(0.0007)	(0.0007)	(0.0006)
Age squared (divided by 1000)	-0.1552	-0.2699	-0.5593	-0.5043
	(0.0301)	(0.0235)	(0.0240)	(0.0234)
Previously full time employed, not on recall:				
Prime aged men (PAM2)				
Being married	0.2657	0.2813	0.2354	0.2334
	(0.0134)	(0.0095)	(0.0096)	(0.0097)
Having small children (up to 5 years)	-0.0062	-0.0336	-0.0225	-0.0431
	(0.0137)	(0.0096)	(0.0094)	(0.0100)
Age (deviation from mean)	-0.0165	-0.0362	-0.0431	-0.0332
	(0.0018)	(0.0013)	(0.0012)	(0.0011)
Age squared (divided by 1000)	0.0548	0.5413	0.8622	0.4239
	(0.1324)	(0.0918)	(0.0832)	(0.0791)
Prime aged women (PAW2)				
Being married	0.0383	0.0993	0.1189	0.1084
	(0.0159)	(0.0111)	(0.0107)	(0.0094)
Having small children (up to 5 years)	-0.2890	-0.2403	-0.3036	-0.3089
	(0.0191)	(0.0129)	(0.0120)	(0.0109)
Age (deviation from mean)	-0.0046	-0.0112	-0.0188	-0.0184
	(0.0014)	(0.0010)	(0.0010)	(0.0009)
Age squared (divided by 1000)	0.0534	0.0163	0.0579	0.1849
	(0.1562)	(0.1065)	(0.0984)	(0.0865)

Note: Standard errors in parentheses.

Figure 5 illustrates how the estimated hazard rate depends on age. The upper left panel displays the time-invariant non-parametric age function estimated with the aid of 50 age dummy-variables. The hazard rate appears to be a relatively smooth and monotonically downwards sloping function of age (conditional on work experience), with a particularly sharp decline after 62 years - the age after which unemployment benefits can be kept indefinitely. The upper right panel presents the period-specific estimates based on quadratic age functions. It is clear that the negative age effect has become more pronounced over time. Previous findings (Røed and Zhang, 1999a) suggest that the increasing age effect has been particularly associated with declining hazard rates for the oldest workers. This is to some extent supported by the estimates reported in the two lower panels. The negative age profile among older workers has become steeper, particularly for women.

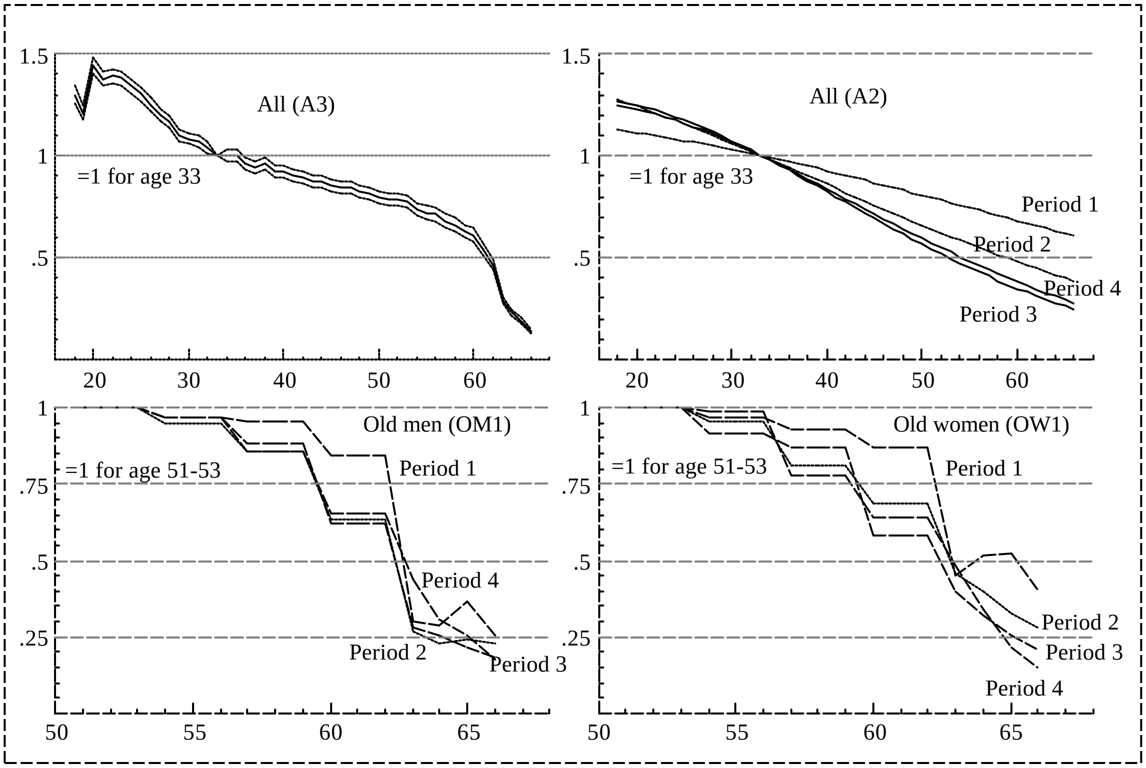

Figure 5. Non parametric (time invariant) and parametric (time varying) age effects
Note: The relative time invariant hazards in the upper left panel are plotted with 95 per cent point-wise confidence intervals. The other panels contain point estimates only. The relative period-specific hazard rates in the upper right panel are calculated on the basis of quadratic age functions, while the relative hazards in the two lower panels are calculated on the basis of piecewise constant age functions covering three-year intervals from 51-63, and then one-year intervals.

5.7 Immigrant Status

Table 9 displays the results regarding the effect of being immigrant from a Non-OECD country. The immigrant hazard rate is substantially below that for Norwegians in all groups. For prime aged male immigrants, the hazard is around 30 per cent lower than for observationally equal Norwegians. There are no significant differences between immigrants with and without Norwegian citizenship. Among the female immigrants, there has been a slight deterioration in the relative hazard for the non-citizens. In the final period the estimated hazard rate is 40 per cent below that for similar Norwegians. The relative hazard rate for immigrants was at it lowest during the initial stages of the recovery (period 3) for all groups, suggesting that employers prefer native employees when they have many job seekers to choose from. Similar results are previously reported by Teulings (1993).

Table 9
Immigrants from Non-OECD Countries

	Period 1	Period 2	Period 3	Period 4
Previously full time employed, not on recall:				
Prime aged men (PAM2)				
Immigrant from a non-OECD country	-0.3576	-0.2550	-0.4045	-0.3426
	(0.0410)	(0.0298)	(0.0259)	(0.0243)
Also Norwegian citizen	0.0316	-0.0339	0.0525	-0.0209
	(0.0659)	(0.0439)	(0.0403)	(0.0294)
Prime aged women (PAW2)				
Immigrant from a non-OECD country	-0.3056	-0.4088	-0.5786	-0.5180
	(0.0601)	(0.0453)	(0.0371)	(0.0299)
Also Norwegian citizen	0.0477	0.1600	0.4178	0.2441
	(0.0908)	(0.0613)	(0.0531)	(0.0393)
Young men and women (YE1)				
Immigrant from a non-OECD country	-0.3378	-0.2306	-0.4107	-0.3679
	(0.0624)	(0.0459)	(0.0464)	(0.0424)
Also Norwegian citizen	0.3016	0.0784	0.2163	0.1826
	(0.0900)	(0.0627)	(0.0617)	(0.0536)
Previously in education:				
Young men and women (YS1)				
Immigrant from a non-OECD country	-0.1976	-0.2230	-0.4833	-0.3779
	(0.0355)	(0.0245)	(0.0242)	(0.0232)
Also Norwegian citizen	0.2035	0.1905	0.3899	0.1725
	(0.0585)	(0.0363)	(0.0349)	(0.0320)
Previously full time employed, on recall				
All (R1)				
Immigrant from a non-OECD country	-0.1807	-0.2533	-0.3184	-0.2818
	(0.0417)	(0.0398)	(0.0385)	(0.0420)
Also Norwegian citizen	0.0723	0.1288	0.0918	-0.0202
	(0.0620)	(0.0549)	(0.0569)	(0.0530)

Note: Standard errors in parentheses.

For youths, there are indications that the negative immigrant-effect has become stronger during the estimation period. The negative immigrant-effect in this group is substantially smaller for immigrants with Norwegian citizenship.

6 Conclusions

We have estimated flexible hazard rate models based on a total number of 1.7 million unemployment spells in Norway 1989-1998, incorporating unrestricted calendar time and duration effects and a large number of person-specific time-varying covariates with time-varying coefficients. Our main conclusions are as follows:

1. The replacement ratio has a significant negative effect on the exit rate out of unemployment. The negative effect is much larger in recessions than in recoveries. The effect is present at all durations, and for some groups it becomes stronger as the spell lengthens.

2. The exit rate out of unemployment increases significantly around the times of transitory and permanent unemployment benefit exhaustion. However, much of the increase occurs just after, rather than just prior to exhaustion.

3. Apart from the spikes associated with benefit exhaustion, there is - conditioned on observed heterogeneity - a strong and monotonic negative duration dependence. But the results are also consistent with exponential duration models mixed with unobserved heterogeneity.

4. The variation in aggregate outflow rates is primarily driven by variation in macroeconomic conditions. During the initial stages of a recession, the fall in the aggregate outflow rate is mitigated by a compositional improvement in the unemployment pool related to the higher inflow rate (and the corresponding larger fraction of short-term unemployed). In the subsequent recovery, the rise in the aggregate outflow rate is dampened by a corresponding compositional deterioration.

5. Long term unemployed and low-skilled are more sensitive towards business cycles than short term unemployed and high-skilled.

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